ECONOMIC GROWTH CENTER

YALE UNIVERSITY

Box 1987, Yale Station
New Haven, Connecticut

CENTER DISCUSSION PAPER NO. 347

# MARKET OPPORTUNITIES, GENETIC ENDOWMENTS AND THE INTRAFAMILY DISTRIBUTION OF RESOURCES: CHILD SURVIVAL IN RURAL INDIA

Mark R. Rosenzweig and T. Paul Schultz

March 1980

Notes: This material is based upon work supported by the National Science Foundation under Grant No. SOC-7814481. Any opinions, findings, and conclusions or recommendations expressed in this publication are those of the authors and do not necessarily reflect the views of the National Science Foundation.

## I. Introduction

The allocation of resources within the family has received much attention by economists in recent years. An important rationale for attempting to understand the determinants of the intra-household distribution of goods and time is that such behavior may importantly influence the outcomes of market-oriented policy interventions which tend to characterize most potential or actual policy programs. For example, the effect of a subvention of educational institutions on the earnings potential of children may depend in part on how parents respond, by reducing or increasing the levels of household "educational" resources devoted to children. Similarly, attempts to equalize the earnings (or earnings opportunities) of men and women may result in reinforcing or counteracting substitution responses on the part of parents with respect to the distribution of household consumption or investment goods to children according to sex; for example, the time of female children may be substituted for that of the mother in household work through reductions in school attendance when adult female wage levels increase (Rosenzweig, 1979). Another important question related to the intrafamily distribution of resources concerns whether parents act so to reduce or augment the effect of differences in genetic endowment on earnings or other measures of achievement (Griliches, 1974). Unfortunately, the measurement of these potentially important phenomena appears to pose severe data requirements, necessitating either extensive household time budget surveys (Evenson, 1978; Butz and DaVanzo, 1978) or the imposition of a large number of unverifiable behavioral assumptions to derive inferences from more conventional data (Lazear and Michael, 1980).

In this paper we examine how intrafamily resource allocations respond to changes in economic conditions and to genetic differences in children by estimating the determinants of variations in the sex-specific survival differential of rural Indian children, based on standard census and household survey data. The theoretical framework utilized yields a number of testable implications concerning household allocation behavior which are obtained without the imposition of special behavioral assumptions or restrictions on the biological determinants of child survival.

While sex differences in child survival or mortality have been examined by social scientists (Ben-Porath and Welch, 1972; Cassen, 1978; El-Badry, 1969; Hammoud, 1977; Ram Gupta, 1975), only one attempt has been made to explore empirically the possibility that the greater mortality rates of girls relative to boys in countries such as India and Pakistan, contrary to the experience in most other places of the world, and the large cross-sectional variations in relative rates in such countries are linked to economic behavior (Bardhan, 1974; Boserup, 1970). We expl in particular the hypothesis that such differences are importantly related to the relative returns to survival, with households selectively allocating resources to children in response to variations in sex-differences in their expected earnings opportunities as adults.

In section II, we formulate a simple model of the household in which the relationships between the allocations of (unobserved) household resources between children, biological differences by sex in the responsiveness of survival propensities to consumption levels, adult sex specific earnings potential and differences in survival rates are set out explicitly. We then derive predictio regarding how changes in differential adult employment and earnings opportunities impact on household resource allocations and on sex differences in survival rates.

In section III we specify the empirical framework to be applied to both 1961 aggregate district-level census data and 1971 household survey data from rural India. Section IV reports the results from the household data, based on household level sex-specific survival rates and in Section V we construct a measure of sex-differences in child survival based on census age distribution information and report results based on district data from all but one state in India. Both levels of analyses indicate that the intra-household allocation of resources is highly responsive to market signals. In particular, it is found, consistent with the model, that where women's expected participation in the labor market or female earnings prospects are relatively high, female childre evidently receive a larger share of household resources relative to male children. Th sex differences in the expected productive roles of adults appear, moreover, to dominate differences in wealth or educational levels in accounting for the variation in the sex-specific survival rates of Indian children.

## II. Theoretical Framework

To examine the relationship between the intra-household allocation of resources and "genetic" (sex) differences in potential earnings and in survival rates and to derive testable implications with respect to the determinants of the latter, we construct a one-period household model with two types of children (male and female). It is assumed that parents derive pecuniary benefits associated with the productive capacities of their children as well as direct consumption benefits from surviving children. The household has a utility function, given by (1)

$$(1) \qquad U = U(x_H, m, f)$$

where $x_H$ is a jointly consumed aggregate consumption good and m and f are the number (health, productivity) of surviving male and female children. Given exogenously fixed levels of male and female births M and F, the household can allocate resources $X_i$, $i = m,f$ to the children to influence their survival. The relationship between the $X_i$ and the number of surviving children is given by sex-specific, biologically determined depreciation functions $\delta_i(X_k)$, such that

$$(2) \qquad m = M(1-\delta_m(X_m)) \qquad \delta_m' < 0$$

$$(3) \qquad f = F(1-\delta_f(X_f)) \qquad \delta_f' < 0$$

If the price index of the goods used to augment survival (health) or reduce depreciation is p and surviving male and female children contribute $R_m$ and $R_f$ respectively to family resources, then the income constraint of the household, ignoring discounting, is

$$(4) \qquad V + R_m(1 - \delta_m(X_m)) + R_f(1 - \delta_f(X_f)) - X_H - p(X_m + X_f) = 0$$

where V is exogenous earnings or income and M and F are arbitrarily set equal to one, so that m and f are survival rates.

The household allocates resources $X_m$, $X_f$ among the children and consumes $x_H$ to maximize (1), subject to (2), (3) and (4). The first-order conditions are:

$$(5) \qquad \frac{U_m}{\lambda} = -\frac{p}{\delta_m'} - R_m$$

$$(6) \qquad \frac{U_f}{\lambda} = -\frac{p}{\delta_f'} - R_f$$

$$(7) \qquad \frac{U_{x_H}}{\lambda} = 1$$

where $\lambda$ is the Lagrangean multiplier.

Expressions (5) and (6) indicate that the shadow price of the sex-specific survival good $X_i$, for given p, is lower the greater is the sex-specfic marginal product of such goods in attenuating mortality and the higher is the sex-specific pecuniary contributions of the surviving child of sex i.

While we are interested in how the household goods $X_i$ are allocated between the two types of children (to M and F), assume that only the survival outco m and f are observed and that, as indicated in (2) and (3), survival rates are in part determined by unobserved biological processes. If we define the difference in <u>levels</u> of sex-specific survival as S, then

(8) $$S = m - f = \delta_f(X_f) - \delta_m(X_m)$$

and it can be seen that the sign of the differential mortality levels cannot indicate unambiguously the relative <u>levels</u> of the household goods allocated among the children, since the magnitudes of the sex-specific depreciation rates $\delta_i$ are also unknown, However, <u>changes</u> in the relative amounts of the $X_i$ can map directly into differences in S, as

(9) $$dS = \delta_f' dX_f - \delta_m' dX_m$$

Expression (9) indicates that some qualitative predictions derived from the model regarding the determinants of variations in S pertain as well to changes in the unobserved distribution of goods within the household, regardless of quantitative differences in the biological responses of male and female survival rates to alterations in the level of resources (the $\delta_i$ functions).

To obtain the effect of a change in the expected earnings returns

$R_i$ on S, and thus on the relative levels of the $X_i$, totally differentiate expressions (4) through (7). Using expression (9) and letting own and cross compensated price effects for m and f be given by $\sigma_{ii}$, $i = m,f$, and $\sigma_{mf}$, and income effects by $\eta_i$, it can be readily shown that

$$(10)\quad \frac{dS}{dR_f} = \delta'_m\sigma_{mf} - \delta'_f\sigma_{ff} + f(\eta_m - \eta_f) = \delta'_m\sigma_{mf} - \delta'_f\,\sigma_{ff} + f\frac{dS}{dV}$$

$$(11)\quad \frac{dS}{dR_m} = \delta'_m\sigma_{mm} - \delta'_f\sigma_{mf} + m(\eta_m - \eta_f) = \delta'_m\sigma_{mm} - \delta'_f\sigma_{mf} + m\frac{dS}{dV}$$

If it is assumed that surviving male and female children are substitutes in (1), $\sigma_{mf} > 0$, and that differences in income effects on surviving children are small, then it can be seen that expressions (10) and (11) are of opposite sign -- a rise in the potential earnings of surviving girls increases their survival rates (consumption of good X) relative to those of boys (decreases S); a rise in $R_m$ similarly increases the survival rates of boys relative to girls (increases S). If, however, surviving girls are a "luxury", as the evidence reported below weakly suggests, it is possible that S will decrease in response to a rise in Rm; S and $R_f$, however, will *a fortiori* be negatively related.

Thus with relatively weak assumptions imposed on the utility function and with *no* assumptions regarding sex-differences in biological propensities to survive as a function of consumption levels, it is possible to derive predictions with respect to how sex-differences in child survival rates and thus the intra household allocation of resources will behave in response to changes in the economic environment. In this case, the model suggests that household resource allocation behavior will *reinforce* market forces which operate to change sex-differences in earnings (consumption) opportunities. Productive capacities are thus differentially augmented by the family in response to differential increases in earnings opportunities as long as surviving male and female children are not strong complements in the utility function.

The relative ease with which predictions regarding the association between relative survival rates and sex-specific earnings can be derived, in the absence of information on biological determinants of survival, is due in large part to the ability to distinguish shadow price components of the family distributional commodities, $X_i$, which uniquely pertain to the survival of each sex. Conversely, to obtain a prediction for the effect of a change in the direct price (p) of the goods used to decrease mortality on S due, for example, to the governmental provision of healthcare or to increases in education levels, requires more restrictions:

$$(12) \qquad \frac{dS}{dp} = \delta_f'\sigma_{ff} - \delta_m'\sigma_{mm} + (\delta_f' - \delta_m')\sigma_{mf} + m\eta_m - f\eta_f$$

It can be seen from (12) that the effect of a rise in the cost of goods used to enhance child survival on <u>relative</u> sex-specific survival rates depends on biological differences in the depreciation functions, on differences in own substitution effects and on differences in income effects. However, since the latter can be directly measured, discrepencies <u>in sign</u> between dS/dp and dS/dV ( $= \eta_m - \eta_f$) provide some indication of the importance of any biological difference in the sex-specific survival functions. Such tests are reported below.

## III. Econometric Framework

In a stable, slowly developing society such as in rural India, parents can reasonably expect that conditions which they face as adults will also condition in a similar way the behavior of their offspring. We thus assume that expectations of the future earnings contributions of children are formed on the basis of contemporaneous sex-specific patterns of adult behavior and that such expectations are considered in the intrafamily allocation of resources. The basic set of equations we estimate, based on the assumption of intertemporal stability, are:

$$(13) \quad R_m = \alpha_{mo} + \alpha_{m1}X_1 + \alpha_{m2}X_2 + \alpha_{m3}X_3 + \alpha_{m4}X_4 + \varepsilon_m$$

$$(14) \quad R_f = \alpha_{fo} + \alpha_{f1}X_1 + \alpha_{f2}X_2 + \alpha_{f3}X_3 + \alpha_{f4}X_4 + \varepsilon_f$$

$$(15) \quad S = \beta_o + \beta_1 R_m + \beta_2 R_f + \beta_3 X_2 + \beta_4 X_3 + \beta_5 X_4 + \mu$$

where $X_1$ is a vector of variables influencing the demand for adult labor services, $X_2$ is a vector of variables which may act to constrain employment — religious proscriptions, other segmentation; $X_3$ is a vector of wealth, production or asset variables and $X_4$ is a vector of variables representing educational attainment. All the $X_j$ variables may influence the earnings contributions of male and female adults differentially.

The error terms in equations (13), (14), and (15) are likely to be correlated as adult employment (one component of $x_H$ in (1)) and child survival are jointly determined out of a utility maximization process and thus reflect the same household preference orderings. Moreover, differences in the relative population sizes of adult males and females, aggregations of past intrafamily allocation decisions by parents, may influence the relative returns of the two groups in the labor market, given imperfect substitution and cost of migration. Equation (15) is thus estimated using predicted values of $R_m$ and $R_f$, based on the parameter estimates from (13) and (14). The specifications assume that the labor demand variables in $X_1$ only influence survival differentials through their effects on the expected earnings contributions variables; all other variables jointly influence $R_m$, $R_f$, and S.

Based on the model, we would expect that $\beta_2 < 0$, $\beta_1 > 0$, with the size of the $\beta_4$ coefficients reflecting the differential income effects on m and f. If it is assumed that schooling contributes to the efficiency with which given resources are combined to augment child survival, i.e., p and schooling are negatively correlated, we would expect that the $\beta_4$ and $\beta_5$ coefficients would display the same signs as long as biological influences in sex-specific survival functions are not important (equation (12)). The signs and magnitudes of the coefficient vectors $\alpha_{i2}$ through $\alpha_{i4}$, i = m,f in (13) and (14) will reflect the influence of both demand and supply effects on adult earnings prospects.

## IV. Empirical Application: Rural Indian Household Data

We first utilize child survival data from a national sample of 4000 rural households from India, collected and coded by the National Council of Applied Economic Research. Households in which the mother was aged 15 to 44 and had borne at least one girl and one boy in districts for which wage data were available were selected for analysis, resulting in a final sample size of 1334 households. We test for the potential bias inherent in this sample selection in section 5 below, where we use aggregate data for almost all districts. The sex ratio at birth in this sample, based on over 5000 births, is 1.08, comparable to the ratio of 1.09 found by Pakrasi and Halder (1971) using the 1961-62 Indian National Sample Survey; both ratios are only slightly higher than the sex ratio at birth in the U.S. of 1.06. Reported sample sex ratios at birth thus do not display any unusual male bias. The survival differential variable for each household was constructed from information on the number and sex of live births, M,F, and the number and sex of children who survived m,f. The measure of S for household k is thus $(m_k/M_k)-(f_k/F_k)$. Table 1 lists the variables used in the analysis and their sample means and standard deviations. As can be seen, the survival differential, which is standardized for the sex ratio at birth, on average favors males.

Two measures of expected differential economic contributions are used. The first, based on sex-specific employment probabilities (on the household's land or in the market) are used as a basis for comparison with the district-level analysis in section V. The second measure, which takes into account the value and quantity of employment, is the employment-weighted earnings difference between adult males and females, $R_m - R_f$, computed from sex-specific district-level agricultural wage rates and individual expected employment rates. We use this measure rather than predicting expected wages separately for each sex, because the distribution of the expected earnings difference more closely approximates normality than does $R_f$, as half of the women in the sample are not earning wages. Both the sex differences in employment rates and the sex differences in wage rates imply, not surprisingly, that the expected earnings power of males is twice that of females in rural India.

Table 2 reports the results of the first stage equations. The set of demand variables, $X_1$, used to identify the second-stage survival differential equations, are as a group statistically significant at the 1 percent level. As expected, in heavily Moslem areas, employment rates of women but not those of men are significantly lower and men earn significantly more on average than do women. Similar results are obtained for areas where lower caste populations dominate. However, while factories in rural areas appear to favor the employment of men, small scale industry,

Table 1



Variable Means, Standard Deviations and Level of Observation, Rural Indian Households, 1971

| Variable | Mean | Standard Deviation | Level of Observation |
|---|---|---|---|
| Endogenous | | | |
| Male-Female Child Survival Differential | .018 | .281 | Couple |
| Female Employment Rate | .479 | .500 | Couple |
| Male Employment Rate | .989 | .105 | Couple |
| Male-Female Expected Earnings Difference (Rupees/Day) | 2.27 | 1.93 | Couple/District |
| Exogenous Included | | | |
| Non-earned Income (Rupees/Year)[1] | 123.9 | 526.1 | Couple |
| Gross Cropped Area (Acres)[1] | 7.81 | 11.1 | Couple |
| No Land (Dummy)[1] | .327 | .469 | Couple |
| Village Electrified (Dummy)[1] | .370 | .483 | Village |
| Normal District Rainfall (mm/Year)[1] | 453.0 | 899.0 | District |
| Wife with Some Formal Education, but Less than Martriculate (Dummy)[2] | .122 | .327 | Couple |
| Head with Some Formal Education, but Less than Matriculate (Dummy)[2] | .564 | .496 | Couple |
| Female Matriculate or above (Dummy)[2] | .023 | .396 | Couple |
| Male Matriculate or above (Dummy)[2] | .194 | .396 | Couple |
| Age of Wife | 39.2 | 8.72 | Couple |
| Age of Husband | 46.0 | 9.75 | Couple |
| Exogenous Excluded | | | |
| Factory in Village (Dummy)[3] | .087 | .281 | Village |
| Small Scale Industry in Village (Dummy)[3] | .095 | .294 | Village |
| District Female Agricultural Wage (Rupees/Day)[3] | 1.95 | .896 | District |
| District Male Agricultural Wage (Rupees/Day)[3] | 3.22 | 1.47 | District |
| Proportion of District Females Moslem, 15-59[4] | .072 | .109 | District |
| Proportion of District Population in Scheduled Castes[4] | .144 | .083 | District |
| Number of Households in Sample | | 1334 | |

[1]Variables treated as $X_3$ including wealth, production and assets.

[2]Variables treated as $X_4$ representing educational attainment.

[3]Variables treated as $X_1$ that only affect the derived demand for adult labor.

[4]Variables treated as $X_2$ that culturally constrain employment.

Table 2

OLS Regression: Prediction Equations, Female and Male Employment Rates and Expected Earnings Difference, Rural Indian Households, 1971

| Variable | Female Employment Rate Coefficient | t | Male Employment Rate Coefficient | t | Expected Earnings Difference Coefficient | t |
|---|---|---|---|---|---|---|
| Non-Earned Income ($x10^{-4}$) | -.416 | 1.69 | -.214 | 3.77 | -.164 | 0.26 |
| Gross Cropped Area ($x10^{-4}$) | -33.8 | 2.61 | 4.60 | 1.49 | 35.0 | 1.14 |
| No Land | .031 | 1.00 | -.019 | 2.58 | -.190 | 2.58 |
| Electrification ($x10^{-2}$) | -1.11 | 0.37 | .562 | 0.80 | -.423 | 0.06 |
| Rainfall ($x10^{-4}$) | .475 | 2.43 | -.055 | 1.20 | -.881 | 1.92 |
| Female Primary Education | -.151 | 3.59 | -.005 | 0.47 | .362 | 3.66 |
| Male Primary Education | -.162 | 5.20 | .011 | 1.50 | .333 | 4.52 |
| Female Matriculate | .199 | 2.23 | -.014 | 0.65 | -.589 | 2.79 |
| Male Matriculate | -.347 | 8.10 | .010 | 1.01 | .623 | 6.15 |
| Female District Wage Rate | -.029 | 1.84 | .010 | 2.78 | -.188 | 5.16 |
| Male District Wage Rate | -.067 | 7.00 | .002 | 0.86 | 1.05 | 4.69 |
| Presence of Factory | -.156 | 3.33 | .015 | 1.38 | .316 | 2.84 |
| Small Scale Industry | .097 | 2.17 | -.028 | 2.62 | -.312 | 2.95 |
| Proportion District Moslem ($x10^{-2}$) | -.781 | 5.88 | -.019 | 0.61 | 1.20 | 3.82 |
| Proportion Scheduled Castes | -.013 | 7.42 | -.001 | 0.10 | .024 | 6.02 |
| Age of Wife ($x10^{-2}$) | -.005 | 0.02 | -.030 | 0.44 | .102 | 0.15 |
| Age of Husband ($x10^{-2}$) | -.109 | 0.42 | -.099 | 1.62 | -.416 | 0.68 |
| Constant | 1.22 | 15.96 | 1.03 | 57.12 | -1.30 | 7.19 |
| $R^2$ | .235 | | .046 | | .713 | |

where present, tends to increase the demand for women relative to men.

Since household behavior can only negligibly affect district wages of men and women, the coefficients on these wage variables can be interpreted as own wage and cross wage effects on the household's demand for member's nonmarket time (Ashenfelter and Heckman, 1974). In accord with empirical evidence from many countries, the wife's employment response to the male wage is negative and substantial in magnitude, the elasticity being -.45. The wife's uncompensated own wage employment response is also negative, but smaller in elasticity terms, -.12. The husband's uncompensated own-wage employment response is essentially zero, whereas his cross-wage elasticity is positive, but small, .03.

Of the other coefficients, with the exception of rainfall, all variables associated with either wealth or income tend to lower the employment probability of women, with income effects tending to reduce the expected earnings of men somewhat more than those of women. Consistent with the hypothesis that women are employed more in wet agriculture, however, there is a significant positive correlation between normal district-level rainfall and the probability that a woman is employed; the correlation is weakly negative for males. Schooling evidently plays an important role in determining the employment rate of women, with female schooling effects non-linear -- primary school women tend to be employed less than women with less education or with matriculate degrees or higher. Higher levels of male schooling are associated with both lower employment rates for women and, weakly, higher rates for men.

The second-stage regression results for the child survival differential are reported in Table 3. The employment rate coefficients are consistent with the reinforcing hypothesis -- a rise in the expected adult male employment rate exacerbates the differential in favor of boys, other things equal. The expected wage difference coefficient, significant at the one percent level, also confirms the positive covariation between the relative economic contributions of adults and the distribution of resources among children

Table 3



Instrumental Variables Regressions

Male-Female Survival Differential, Rural Indian Households, 1971

| Variable | Coefficient | t | Coefficient | t |
|---|---|---|---|---|
| Female Employment Rate[a] | -.102 | 1.95 | - | - |
| Male Employment Rate[a] | .317 | 0.46 | - | - |
| Expected Earnings Difference[a] | - | - | .015 | 2.80 |
| Non-Earned Income ($x10^{-4}$) | -.031 | 0.14 | -.084 | 0.54 |
| No Land | .053 | 2.20 | .043 | 2.19 |
| Gross Cropped Area ($x10^{-4}$) | -.123 | 0.13 | -.945 | 0.11 |
| Electrification ($x10^{-2}$) | -.818 | 0.43 | -1.06 | 0.59 |
| Rainfall ($x10^{-4}$) | -.154 | 1.46 | -.200 | 1.96 |
| Female Primary Education | -.008 | 0.30 | -.002 | 0.07 |
| Male Primary Education | -.036 | 1.57 | -.025 | 1.23 |
| Female Matriculate | .023 | 0.38 | .010 | 0.17 |
| Male Matriculate | -.075 | 2.21 | -.053 | 1.93 |
| Age of Wife ($x10^{-2}$) | .001 | 0.64 | .001 | 0.41 |
| Age of Head of Household ($x10^{-2}$) | .001 | 0.57 | .001 | 0.84 |
| Constant | .311 | 0.42 | -.085 | 2.07 |

[a]Endogenous variable

within the household. The point estimates indicate that a rise in the adult female employment rate to .66, an increase of 37 percent, would erase the mean survival differential, a one rupee decrease in the expected daily earnings differential between men and women would similarly eliminate the imbalance in sex-specific survival rates. As indicated however, parity in survival rates does not necessarily reflect parity in the distribution of household goods.

Holding constant the relative expected employment rates, the results suggest also that increases in wealth, in terms of land or other productive capital and asset income,are associated with greater survival prospects of female children. As a consequence of the apparent "superiority" of female survival, boys appear to have significantly higher survival rates relative to girls in landless than in landed households. The negative signs for three of the four schooling variables (the fourth is not statistically significant) thus suggest that biological differences in survival propensities between the sexes may be unimportant, as education and income effects appear qualitatively similar.

V. Empirical Application: Indian District-level Data

While aggregate and regional (state) data from India on child and infant mortality rates also appear to document significant variations and anomalies in sex differences in child survival (Bardhan, 1974), it is well-known that vital rate data from India are of insufffficient quality to withstand an intensive analysis of these phenomena at a finer geographical level. Indian census age distribution data, however, appear to accurately capture sex-differences in population size by age (Visaria, 1969). In this section we show that such data can be used to depict sex-specific differences in survival rates and thus provide information on intra-family allocation behavior. Based on the constructed measure, we than apply the model to rural district-level data assembled from the 1961 Indian census for 295 districts (those which report rainfall data), representing more than 95 percent of the total rural Indian population to analyze the determinants of the variation in sex-specific child mortality differences across districts.

To derive the relationship between sex-specific age structure and mortality differentials assume that the population is stable (Coale, 1972). The sex ratio of the number of males to females aged a in a stable population is:

$$(16) \qquad \frac{\bar{C}_m(a)}{\bar{C}_f(a)} = \frac{b_m e^{-r_m a} p_m(a)\bar{P}_m}{b_f e^{-r_f a} p_f(a)\bar{P}_f}$$

where $b_i$ is the birth rate, $r_i$ is the annual rate of increase, $p_i(a)$ is the proportion surviving from birth to age a, and $\bar{P}_i$ is the total number of persons of the i'th sex in the population and $\bar{C}_i(a) = c_i(a)\bar{P}_i$, and c(a) is the proportion age a.

Rearranging terms, we obtain:

$$(17) \qquad \frac{b_m\bar{P}_m}{b_f\bar{P}_f} \qquad e^{(r_f-r_m)a} \qquad \frac{P_m(a)}{p_f(a)}$$

where the first term is the sex ratio at birth, the second term is the differential population growth rates of females to males over "a" years, and the third term is the sex ratio of survival rates. Constructing a ratio of the sex composition of two adjoining age groups t years apart yields:

$$(18) \qquad R(a+t,a) = \frac{\;\frac{\bar{C}_m(a+t)}{\bar{C}_f(a+t)}\;}{\;\frac{\bar{C}_m(a)}{\bar{C}_f(a)}\;} = e^{(r_f-r_m)t} \; \frac{\;\frac{p_m(a+t)}{p_f(a+t)}\;}{\;\frac{p_m(a)}{p_f(a)}\;}$$

or, since $p_i(a) = e^{-\int_0^a \mu_i(x)dx}$

$$R(a+t,a) = e^{(r_f-r_m)t} e^{\int_a^{a+t} \mu_f(x) - \mu_m(x)\, dx}$$

where $\mu_i(x)$ is the sex-specific death rate at age x.

Taking natural logarithms,

$$\text{(19)} \qquad \ln R(a+t,a) = (r_f-r_m)t + \int_a^{a+t} \mu_f(x) - \mu_m(x)\, dx,$$

it can be seen that the logarithm of the ratio of the sex ratios of adjacent age groups in a (sex specific) stable population is a function of possibly distinct sex specific rates of growth of the population and the difference between the survival rates of females and males in the relevant age interval. Since the former is likely to be negligible in most populations even when survival rates generally favor males, as in India, expression (14), the log of adjacent age-group sex ratios measures sex differences in mortality rates over the relevant age range.[1]

Table 4 reports the variables used in the district-level analysis and their sample statistics. The mean child survival difference is .045, and the logarithm of the ratio measure of this variable suggested by the stable population model is .042. According to the above derivation, boys thus appear to have had a four percent greater survival rate than did girls in 1961, somewhat higher than the 2 percent differential observed in the 1971 household sample. The census district data also imply similar rates for male and female adult employment as those obtained from the household survey, with female employment rates somewhat higher in the more recent period and male employment slightly lower. This is a common pattern of post-war change in sex-specific participation rates observed in low and high income countries (Durand, 1975).

Because wage rates are only available for a limited number of districts, only the adult employment rates are analyzed at the district-level. Note that it would not be appropriate to utilize district wage rates as demand variables that identify the employment equations, as was assumed at the household level, since at the aggregate level both employment rates and wages are jointly determined. The list of identifying variables also excludes religion and caste, which at the district level seem more likely to affect directly both adult labor market behavior and sex specific child survival. This exclusion will be reconsidered below. The remaining variables are defined analogously at the district level as they were at the household level, whenever data permit parallel specification. The proportion of the district population living in rural areas is added, however, to represent urban influences.

Table 4

Variable Means and Standard Deviations, Indian Districts, Rural Population, 1961

| Variable | Mean | Standard Deviation |
|---|---|---|
| Endogenous | | |
| Male-Female Child Survival Differential | .0391 | .0446 |
| Log Male-Female Child Survival Ratio | .0415 | .0471 |
| Female Employment Rate, 15-59 | .548 | .257 |
| Male Employment Rate, 15-59 | .943 | .0305 |
| Exogenous Included | | |
| Average Farm Size (Acres)[1] | 12.64 | 13.86 |
| Percentage Households with No Land[1] | 30.61 | 13.98 |
| Proportion of Land Irrigated[1] | .194 | .185 |
| Normal Rainfall (mm/year)[1] | 113.8 | 59.09 |
| Proportion of District Population Rural[1] | .846 | .114 |
| Proportion Females with Primary Education, 15-59[2] | .0259 | .0349 |
| Proportion Males with Primary Education, 15-59[2] | .113 | .0825 |
| Proportion Females Matriculate , 15-59[2] | .265 | .646 |
| Proportion Males Matriculate , 15-59[2] | .0267 | .0231 |
| Proportion of Females Moslem, 15-59[3] | .0709 | .0857 |
| Proportion of Population in Scheduled Castes[3] | .163 | .098 |
| Exogenous Excluded | | |
| Number of Factories per Household[4] | .148 | .184 |
| Percentage of Factories with 5+ Employees[4] | 4.05 | 4.50 |
| Percentage of Factories Using Fuel[4] | 25.12 | 22.37 |
| Number of Districts | 295 | |

[1]Variable treated as $X_3$ including wealth, production or assets.

[2]Variables treated as $X_4$ representing educational attainment.

[3]Variables treated in $X_2$ that culturally constrain employment.

[4]Variables treated in $X_1$ that only affect the derived demand for adult labor.

The estimated prediction equations for female and male employment rates are presented in Table 5 and appear similar to those obtained from the 1971 household data. The number of factories in the district is associated with substantially greater female employment and somewhat greater male employment rates, whereas the share of large factories with five or more employees is again correlated with lower levels of male employment. The only indicator in the census of the technology of the production units is whether they use fuel, which appears unrelated to employment of men or women.

Moslem districts again reveal significantly lower female employment rates, as do those districts with a greater proportion of the population in scheduled castes. In the district sample, rainfall is no longer correlated with women's employment, but instead the share of irrigated land is inversely, and farm size is directly associated with women's employment. District level primary and matriculate education of men and matriculate education of women is correlated with lower male employment, reflecting perhaps both the substitution of time from labor market activities to school work at younger ages, and a wealth effect on the demand for leisure at later ages.

The instrumental variable regressions for rural Indian districts are reported in Table 6, with and without the inclusion of the religion and caste variables. As in the micro-data, higher female employment is associated with a lower male to female survival ratio, significant at the 5 percent level. As before, the male employment coefficient is positive and not significantly different from zero. These results suggest that the micro estimates were not mainly due to sample selection procedure. According to the first regression including the Moslem and caste variables, an increase in district female employment by one-half, from 55 to 83 percent, would lower the male-female survival ratio from 1.04 to unity.

Table 5

OLS Regression: Prediction Equations,

Female and Male Employment Rates (Percent), Rural Indian Districts, 1961

| Variable | Female Coefficient | Female t | Male Coefficient | Male t |
|---|---|---|---|---|
| Mean Farm Size | .203 | 1.99 | .0164 | 1.56 |
| Percent of Households with No Land | -.0548 | 0.46 | -.0186 | 1.52 |
| Percent Irrigated Land | -.334 | 4.07 | -.0044 | 0.53 |
| Rainfall ($x10^{-2}$) | .0236 | 0.84 | -.0001 | 0.03 |
| Proportion District Rural | 12.63 | 1.15 | -.757 | 0.67 |
| Percent Female Primary Education | -.358 | 0.44 | .140 | 1.67 |
| Percent Male Primary Education | -.112 | 0.42 | -.132 | 4.89 |
| Percent Female Matriculate | -2.51 | 0.82 | -2.08 | 6.58 |
| Percent Male Matriculate | -.218 | 0.34 | -.212 | 3.21 |
| Proportion Females Moslem | -121.2 | 8.42 | -2.10 | 1.42 |
| Proportion in Scheduled Castes | -57.39 | 4.41 | -2.75 | 2.07 |
| Number of Factories per Household | 25.39 | 3.72 | 1.25 | 1.78 |
| Percent Factories with 5+ Employees | -.315 | 1.07 | -.135 | 4.50 |
| Percent Factories Using Fuel | -.0452 | 0.79 | -.0045 | 0.76 |
| Constant | 67.38 | | 98.76 | |
| $R^2$ | .488 | | .617 | |

In districts in which a greater proportion of the households are landless, female child survival relative to male is higher, confirming the household-level finding. Irrigation and rainfall are only weakly associated with survival prospects for girls relative to boys, representing either the enhanced productivity of female labor in wet agricultural crops (Bardhan, 1974), or the effect of increased household wealth for a given farm size in regions with more ample water supplies. There is also an indication that urban influences in the district are associated with increased female survival relative to male, a pattern noted by other authors studying the expectation of life at birth (Preston and Weed, 1976). The coefficients on the education variables are negative in sign, consistent with their representing wealth, but they are not statistically important.

A notable finding from these estimates is that districts in which larger proportion of the population is Muslim do not exhibit distinctly different sex specific child survival rates, holding constant for predicted women's employment rates. Indeed, the religion and caste variables are not jointly significant in the first regression, at even the 50 percent level. In the second regression in Table 6 the religion and caste variables are excluded, under the assumption that they affect the intra family allocation of resources to boys and girls only through their effect on adult employment rates for men and women. In this second specification, the female employment rate coefficient is smaller, but its standard error declines by two-thirds. The precision of the estimates of the effects of landless households, irrigation and rainfall is also increased substantially.

Table 6

Instrumental Variables Regression: Log Male-Female Child Survival Ratio; Rural Indian Districts, 1961 Census

| Variable | Coefficient | t | Coefficient | t |
|---|---|---|---|---|
| Female Employment Rate[a] | -.152 | 1.95 | -.111 | 4.14 |
| Male Employment Rate[a] | .204 | 0.33 | .0161 | 0.03 |
| Mean Farm Size ($x10^{-3}$) | .158 | 0.64 | .103 | 0.44 |
| Percentage with No Land ($x10^{-3}$) | -.673 | 2.50 | -.680 | 2.70 |
| Proportion Irrigated Land ($x10^{-3}$) | -.445 | 1.45 | -.333 | 1.75 |
| Rainfall ($x10^{-4}$) | -.693 | 0.99 | -.849 | 1.37 |
| Proportion District Rural | .041 | 1.46 | .035 | 1.39 |
| Proportion Female Primary Education ($x10^{-2}$) | -.108 | 0.49 | -.048 | 0.24 |
| Proportion Male Primary Education ($x10^{-4}$) | -1.98 | 0.21 | -1.42 | .050 |
| Proportion Female Matriculate ($x10^{-2}$) | -.229 | 0.16 | -.547 | 0.44 |
| Proportion Male Matriculate ($x10^{-2}$) | .144 | 0.75 | .114 | 0.63 |
| Proportion Females Moslem ($x10^{-1}$) | -.486 | 0.51 | - | - |
| Proportion of Population in Scheduled Castes | -.030 | 0.70 | - | - |
| Constant | -.057 | | .098 | |

[a]Endogenous variable

## VI. Conclusions

A central working assumption underlying the economic literature on household behavior is that a family utility function exists permitting study of intergenerational allocation as an orderly optimizing process. A question for study is, therefore, whether parents allocate their investments across offspring in order to complement the distribution of genetic endowments of their children or to compensate for these endowments and thereby to equalize the economic opportunities available to their children . In a low income country child survival may be a sensitive indicator of parent investment in children. Our analysis of data from households and districts of rural India suggests that the preference for compensating investments in offspring, assumed, for example, in Becker and Tomes (1976), does not dominate behavior, but rather the random differences in genetic traits of children, associated with sex in this case, evoke reinforcing allocations of family resources.[2] Those children expected to be more economically productive adults receive a larger share of family resources and have a greater propensity to survive. These results imply that attempts to equalize the earnings opportunities of men and women in the current generation therefore may reduce the dispersion of earnings <u>more</u> in future generations than in the contemporaneous period. However, the general association indicated between household wealth and improved female relative to male survival may also imply that at least greater equality in survival opportunities between children is sought as the wealth of families increases, holding constant for the market social valuation of the genetic endowment of sex.

# Footnotes

[1]The closeness of the relationship between the log age-ratio measure and the variable of interest, the mortality differential, depends on how closely the stable population assumption corresponds to the actual age-data and the accuracy of the survey information on the population. The typical problem in using the stable population model for study of demographic processes is that the assumption of past constancy of the fertility and mortality schedule with respect to age is not valid. In this case, a change over time in fertility presents no problem unless it implies a change in the sex ratio at birth or the sex ratio of child mortality. The former appears unlikely and it is the latter which is the focus of the analysis. Moreover, while in analyses dealing with small regional populations, the stable population model is rarely appropriate because of internal migration, where attention is focused only on the sex composition of children less than age 10 there is less reason to expect migration would be sufficiently sex and age selective as to distort interregional comparisons. There is also the possibility that underenumeration of children differs from region to region and perhaps embodies a non-uniform error for boys and girls. Again, however, the construction of the age-ratio form of the measure of child mortality is such that sex-specific errors that persisted from one age group to the next would be offset and any systematic omission of very young or older children would not imply a problem unless it were more common for one sex. On the accuracy of the Census sex/age data, see Visaria (1969) and Natarajan (1972).

[2]The tendency for household investments to reinforce adult market productivities is also evidenced in rural Indian child schooling and employment data. District adult female wages have twice the positive effect on the

school attendance rates of girls than they do of boys, and conversely twice the deterrent effect on the employment rates of girls than of boys. In contrast, adult male wages depress child schooling and employment rates of female children by more than they do male children (Rosenzweig and Evenson, 1977).

# Appendix

District-Level Data Sources: agricultural wage rates, India Directorate of Economics and Statistics (1976); normal rainfall, irrigation, farm size, India, Directorate of Economics and Statistics (1970). All other data, India, Office of the Registrar General (1965): age distribution, Part II-A; religion, caste, schooling, Part II-B; employment, Part II-C; factories, Part IV-B.

Individual Household Data Source: described by Sarma (1975), and available from NCEAR, New Delhi, India.

# References


Ashenfelter, O and J.J. Heckman. January 1974. "The Estimation of Income and Substitution Effects in a Model of Family Labor Supply." Econometrica vol. 42. pp. 73-85.

Bardhan, P.K. 1974. "On Life and Death Questions." Economic and Political Weekly (New Delhi) 9 (32-34), Special Number (August). pp. 1293-1305.

Becker, G. and N. Tomes. August 1976. "Child Endowments and the Quantity and Quality of Children." Journal of Political Economy vol. 84, pt. 2.

Ben-Porath, Y. and F. Welch. 1972. Chance, Child Traits and Choice of Family Size , R-1117. Santa Monica, CA.: Rand Corporation.

Boserup, E. 1970. Woman's Role in Economic Development. New York: St. Martin's Press.

Butz, W.P. and J. DaVanzo. April 1978. Contracepting, Breastfeeding and Birthspacing in Malaysia RCN-6166.

Cassen, R.H. 1978. India: Population, Econcmy and Society. London: Macmillan Press, Ltd.

Coale, A.J. 1972. The Growth and Structure of Human Populations. Princeton, New Jersey: Princeton University Press.

Durand, J.D. 1975. The Labor Force in Economic Development. Princeton, New Jersey: Princeton University Press.

El Badry, J.M. 1969. "Higher Female Than Male Mortality in Some Countries of South Asia." Journal of the American Statistical Association, vol. 64. pp. 1234-44.

Evenson, R.E. May 1978. "Time Allocation in Rural Philippine Households." American Journal of Agricultural Economics, vol. 60, No. 2. pp. 322-330.

Griliches, Z. 1974. "Comment on Toward a New Theory of Population and Economic Growth." In T.W. Schultz (ed.), Economics of the Familv. Chicago: University of Chicago Press.

Hammoud, E.I. 1977. "Sex Differentials in Mortality." World Health Statistics Report 30 (3): 174-206.

India, Directorate of Economics and Statistics. 1976. Agricultural Wages in India 1970-71. New Delhi.

______. Office of the Registrar General. 1965. Census of India, 1961, Parts II-B, II-C, IV-B. New Delhi.

______, Directorate of Economics and Statistics. 1970. Indian Agricultural Statistics, 1961-62 and 1962-63, Vol. 11. New Delhi.

Lazear, E. P. and R.T. Michael. May 1980. "Real Income Equivalence Among One-Earner and Two Earner Families." American Economic Review.

Preston, S.H. andJ.A. Weed. 1976. "Causes of Death Responsible for International and Intertemporal Variation in Sex Mortality Differentials." World Health Statistics Reports 30 (4): 144-188.

Ram Gupta, San T. 1975. "Mortality Patterns in India." In R.B. Chari (ed.) Demographic Trends in India. New Delhi: D.P. Sharma and R.K. Bhatia, pp. 43-52.

Rosenzweig, M.R. Oct. 1979. "Household and Non-Household Activities of Youths: Issues of Modelling, Data Collection, and Estimation." Paper presented at ILO Workshop on Child Employment. Geneva.

______ and R.E. Evenson. 1977. "Fertility, Schooling and the Economic Contribution of Children in Rural India." Econometrica 45 (3): 1065-1080.

Sarma, M.T.R., et al. 1975. Changes in Rural Income in India. New Delhi, India: National Council of Applied Economic Research.

Visaria, P.M. 1969. The Sex Ratio of the Population of India. Census of India 1961, Vol. I, Monograph 10. New Delhi, India: Office of the Registrar General, ud.